# Assessing Intergenerational Earnings Persistence Among German Workers*

*Philipp Eisenhauer and Friedhelm Pfeiffer***

"The vitality and stability of our democracy – as well as the economy – eventually depend on the social permeability of our society."
(Horst Köhler, German Federal President, 29. 12. 2007, authors' own translation)

This statement draws attention to the strong meritocratic beliefs concerning the equality of opportunity that dominate public debates. This is especially true of the education system. But does this general concern translate into a society in which one's economic success in the labor market is independent of the family into which one was born? And if so, to what degree?

In this study, we investigate intergenerational earnings persistence among German workers. Our measure of labor market success is real monthly earnings before taxes and social security contributions. The relationship between fathers' and sons' labor market earnings is assessed using samples drawn from the German Socio-Economic Panel (SOEP) 1984–2006. We introduce a novel sampling procedure that allows us to observe father-son pairs at a fairly similar stage in their lives.

From a variety of microeconometric estimates (utilizing both OLS and IV methods) we suggest that the best point estimate of intergenerational earnings elasticity among German workers is one-third. Hence, if in the period of investigation a father's permanent labor market earnings increased by 10 percent (€ 231 at the mean of our father sample), the son's long-run economic status grew by 3.33 percent. Evaluated at the mean of our sample of sons (€ 1,937), this implies a step up of € 63 for the son.

This figure indicates a lower degree of mobility (and a higher degree of persistence) in Germany compared to preceding studies. In an international perspective, the intergenerational earnings persistence in Germany seems to be lower than that in the United States and higher than that in Sweden. To summarize: there still seems to be substantial intergenerational earnings mobility among German workers, but more persistence than previous research suggested.

* This paper was released for publication in April 2008.
** We thank the Leibniz Association for supporting this study in the research network Noncognitive Skills: Acquisition and Economic Consequences. Furthermore, Friedhelm Pfeiffer's research was supported by the German Science Foundation's grants PF 331/2 (Microeconometric Methods to Assess Heterogeneous Returns to Education) and PF 331/3 (Wages, Rent-Sharing and Collective Wage Bargaining). For helpful discussions we thank Johannes Gernandt, Maresa Sprietsma, Thorsten Vogel, participants of the 7th meeting of the DFG Workshop "Flexibility in Heterogeneous Labour Markets" on 4th October 2007 in Nuremberg, two anonymous referees, and our colleagues at the ZEW. We thank Alina Botezat, Moritz Meyer and Fabian Kosse for their fine research assistance. Any remaining errors are our own.

## Contents

## 1 Introduction and motivation

The vitality and stability of our democracy – as well as the economy – eventually depend on the social permeability of our society.[1]

This statement draws attention to the strong meritocratic beliefs concerning the equality of opportunity that dominate public debates. This is especially true of the education system. According to public rhetoric the aim is to guarantee social mobility in Germany. Families receive a child benefit transfer, schooling for up to 13 years is free of charge and, if education is continued at a university, the cost of living is covered by federal aid for students from low-income families. But, does this general concern translate into a society in which one's economic success is independent of the family into which one was born? And if so, to what degree?

To empirically analyze the intergenerational relationship, the following econometric model

$$y_1^i = \alpha + \beta_y y_0^i + \epsilon_1^i \qquad (1)$$

is used as a starting point (Corak 2004). A linear relationship between the long-run economic status $y_0^i$ and $y_1^i$ of family $i$ in generation 0 and 1 is assumed, allowing for shifts in the mean economic status irrespective of parental status via the parameter $\alpha$. Deviations from the predicted status due to market luck or other random elements in the intergenerational transmission of skills and personal traits are summarized in the idiosyncratic error term $\epsilon_1^i$. Ideally, permanent earnings are chosen as the measure of economic status (Friedman 1957). We use both terms to describe the long-run economic success of an individual. In the case all status variables are measured in their natural logarithm, $\beta_y$ in equation (1) is the intergenerational elasticity of permanent earnings. It measures the (expected) percentage change in the offspring's economic status associated with a one percent change in their parental success. In principle, $\beta_y$ can take any value but most studies find a value between zero and one.[2] A positive value indicates generational persistence of permanent earnings in which higher parental long-run status favors the economic success of the offspring; a negative figure indicates generational reversal of economic status. A value of zero for the intergenerational elasticity $\beta_y$ (child's and parents' economic success are unrelated) corresponds to complete intergenerational mobility, while a value of unity (the child's economic success is completely determined by the parental achievement) is associated with complete immobility. $(1 - \beta_y)$ provides a measure of the degree to which economic status regresses to the mean (Becker & Tomes 1986; Goldberger 1989). If it takes the value one ($\beta_y = 0$), a child of parents who attain a below average long-run status can expect an average status just like the offspring of high-status parents.

Although there is agreement about the existence of an intergenerational link in economic status, a number of recent studies debate its varying magnitude across countries (Solon 2002; Grawe 2006; Jäntti, Røed, Naylor, Björklund, Bratsberg, Raaum, Österbacka & Erikson 2006; Vogel 2007). While many features of the human skill formation process are universal, there may however be unique features in German data. In an international perspective, low tuition fees and federal student aid might ease the impact of borrowing constraints and thus enhance mobility in Germany compared to other countries.

The contribution of our paper to the literature on intergenerational persistence is twofold. First, based upon recent improvements in the understanding of the association between short- and long-run economic status we assess the potential biases in previous studies. Deviations of current from permanent economic status arise due to transitory fluctuations (Bowles 1972; Solon 1992) and a time-varying association between the two (Haider & Solon 2006; Grawe 2006). We introduce a novel sampling procedure that takes both into account and makes it possible to observe father-son pairs at a fairly similar stage in their lives. Second, the relationship is assessed for Germany with samples drawn from the German Socio-Economic Panel (SOEP) 1984–2006.

Our results suggest that the best conservative point estimate of intergenerational earnings persistence among western German workers is 1/3. This indicates a lower degree of mobility (and a higher degree of persistence) in Germany compared to Couch & Dunn (1997) and Wiegand (1997) but is in line with Vogel (2007), who compares intergenerational mobility in Germany and the United States. In an international perspective, the intergenerational earnings persistence seems to be lower compared to the United States $\beta_y^{US} = 0.4$ (Solon 1992), and higher compared to Sweden $\beta_y^S = 0.2$ (Björklund & Jäntti 1997). There still seems to be substantial intergenerational earnings mobility among western German workers, but more persistence than previous research suggested.

[1] Horst Köhler, German Federal President, in an interview with the Frankfurter Allgemeine Zeitung, 29. 12. 2007, Berlin (authors' own translation).
[2] See Solon (2002) for a recent survey.

The remainder of this paper is organized as follows: Section 2 provides an introduction to the econometric methods applied to estimate intergenerational persistence with incomplete data. Section 3 presents our novel sampling procedure with the SOEP. Section 4 discusses the econometric findings, and section 5 concludes.

## 2 Econometric problems and findings from the literature

In this section the econometric problems associated with measuring intergenerational persistence and the conclusions we draw regarding its estimation among German workers are pointed out.

### 2.1 Measurement error problems

The deduction of an individual's permanent earnings requires a lifelong earnings history. Since researchers usually lack direct measures of long-run status $y_0^i$ and $y_1^i$ for two generations in order to investigate intergenerational mobility, they rely on proxies ($y_{0h}^i$, $y_{1t}^i$) of permanent earnings for each generation (0,1) observed at ages $h$ and $t$. Sometimes only single-year measures of earnings[3] have been used. Usually, however, a short-run measure of economic status is an imperfect proxy of long-run status. It is subject to measurement error due to transitory fluctuations and lifecycle variation in the association between current and lifetime earnings.[4]

#### 2.1.1 Transitory fluctuations

Current earnings of fathers $y_{0h}^i$ and sons $y_{1t}^i$ can be decomposed as follows (Friedman 1957).

$$y_{1t}^i = y_1^i + v_{1t}^i \tag{2}$$

$$y_{0h}^i = y_0^i + v_{0h}^i \tag{3}$$

$y_0^i$ and $y_1^i$ describe time-invariant permanent earnings, while ($v_{0h}^i$, $v_{1t}^i$) indicates time-varying transitory fluctuations. The latter might arise from job mobility, business cycle effects or variable compensation schemes. If current earnings deviate from the permanent status, using them as a proxy for the long-run status introduces attenuation bias in the estimation of equation (1). Assuming that $v_{1t}^i$ and $v_{0h}^i$ are uncorrelated with each other and permanent earnings $y_0^i$ and $y_1^i$, a deviation implies a downward inconsistency of the estimated slope coefficient $\hat{\beta}_y^{OLS}$ in an OLS estimation by the factor $\theta_h$ (Solon 1992).

$$\text{plim } \hat{\beta}_y^{OLS} = \theta_h \beta_y < \beta_y \tag{4}$$

$$\theta_h = \left( \frac{Var[y_0]}{Var[y_0] + Var[v_{0h}]} \right) \tag{5}$$

The attenuation factor $\theta_h$ captures how much signal $Var[y_0]$ is provided by the measure $y_{0h}$ relative to its total noise, $Var[y_{0h}] = Var[y_0] + Var[v_{0h}]$.

Based on single-year snapshots, empirical findings by Corcoran, Laren, Gordon & Solon (1991), Card (1994) and Hyslop (2001) suggest an attenuation factor around $\theta_h = 0.5$. This implies a (considerable) signal-to-noise ratio of observed parental earnings and an attenuation bias of $(1 - \theta_h) = 0.5$. Note also, that transitory fluctuations in the offspring's earnings $v_{1t}^i$ do not bias the OLS estimation in equation (1) as long as they are uncorrelated with $v_{0h}^i$. However, the higher their variance, the larger the confidence interval of $\hat{\beta}_y^{OLS}$ will be.

*Averaging parental earnings*

To decrease the magnitude of the inconsistency, Solon (1992) suggests averaging parental status over $T$ years, which reduces the variance of the noise relative to the signal. Transitory shocks are averaged away as long as the process is stationary, see Mazumder (2005).

$$\theta_h = \left( \frac{Var[y_0]}{Var[y_0] + \frac{1}{T} Var[v_{0h}]} \right) \tag{6}$$

As more years of data are used, the attenuation factor $\theta_h$ rises and the attenuation bias $(1 - \theta_h)$ declines. According to Mazumder (2005), the attenuation factor $\theta_h$ rises to $\theta_h = 0.7$ (from $\theta_h = 0.5$) when relying on a 5-year average of earnings. The attenuation bias is reduced to $[(1 - \theta_h) = 0.3]$. Solon (1992) and Wiegand (1997) estimated an intergenerational elasticity of fathers' and sons' earnings based on 5-year averages of 0.4 for the United States and 0.2 for Germany. Given the attenuation factor mentioned above, the true elasticities would come closer to 0.6 for the United States and 0.3 for Germany.

[3] See Behrman & Taubman (1985) as an example.
[4] For a further errors-in-reporting problem see Bound & Krueger (1991) and Duncan & Hill (1985).

### Instrumenting parental earnings

In a second approach to estimate $\beta_y$, the direct projection of $y_1^i$ on $y_0^i$, Solon (1992) proposes an IV estimation. Acknowledging the difficulties of finding an instrument that is correlated with parental long-run status but not a structural determinant of the offspring's permanent earnings, the basic model in equation (1) is amended to include an additional factor $I_0^i$.

$$y_1^i = \beta_1 y_0^i + \beta_I I_0^i + \omega_1^i \tag{7}$$

In this case, performing an IV estimation of $\beta_y$ using $I_1^i$ as the instrument yields the following probability limit, Solon (1992).

$$\text{plim } \hat{\beta}_y^{IV} = \beta_y + \beta_I \left(\frac{1-\varkappa^2}{\varkappa}\right)\left(\frac{Sd[I_0]}{Sd[y_0]}\right) \tag{8}$$

$$\varkappa = \frac{Cov[I_0, y_0]}{Sd[y_0]Sd[I_0]} \tag{9}$$

$\hat{\beta}_y^{IV}$ is an unbiased estimator for $\beta_y$ only if the instrument does not influence the offspring's status ($\beta_I = 0$) or the instrument and parental status are perfectly correlated, $|\varkappa| = 1$. The closer $|\varkappa|$ is to one, the smaller the bias as there is less variation in earnings that is not captured by the instrument. Assuming a positive but imperfect correlation between the instrument and parental long-run status, the direction of the inconsistency is determined by $\beta_I$. If the instrument $I_0^i$ has a positive impact on the offspring's status ($\beta_I > 0$), the estimator will be biased upward. If the opposite is true, the estimated coefficient is downward biased like the OLS estimate.

In empirical research, parental years of education (Solon 1992; Dearden, Machin & Reed 1997) or indicators of occupational prestige (Zimmerman 1992; Wiegand 1997) have been used to instrument long-run parental status. Since years of education enhance labor market earnings, it may capture an important part of parental permanent earnings, although not necessarily 100 %.[5] In this case an IV estimate using years of education will be upward biased.

Estimating the intergenerational elasticity $\hat{\beta}_y$ using OLS and IV techniques therefore suggests bracketing the coefficient (Solon 1992). The OLS estimate is downward inconsistent due to error-in-variable bias, whereas the IV estimate is presumably upward biased. Accounting for the associated standard errors, $\beta_y$ is located between the two estimates.

$$\boxed{\hat{\beta}_y^{OLS} < \beta_y < \hat{\beta}_y^{IV}}$$

### 2.1.2 Lifecycle variations

Empirical research as well as theoretical reasoning suggest that wage workers differ with respect to their age-earnings profiles.[6] This may occur due to age-specific heterogeneity in human capital investment or variations in the wage structure across jobs set up by firms for the purpose of effort regulation and incentive compatibility. For estimation purposes, the projection of current on permanent earnings is generalized to include a time-varying parameter $\lambda_{t,h}$ to capture age-specific aspects in the association between current and permanent earnings over the lifecycle (Haider & Solon 2006).

$$y_{1t}^i = \lambda_t y_1^i + v_{1t}^i \tag{10}$$

$$y_{0h}^i = \lambda_h y_0^i + v_{0h}^i \tag{11}$$

Averaging parental earnings $y_{0h}^i$ across $T$ years, the interaction of both types of measurement error is considered. If the parents' and the offspring's long-run status is proxied by short-run earnings, equation 12 determines the potential bias.

$$\text{plim } \hat{\beta}_y^{OLS} = \lambda_t \theta_h \beta_y \tag{12}$$

$$\theta_h = \frac{\lambda_h Var[y_0]}{\lambda_h^2 Var[y_0] + \frac{1}{T} Var[v_{0h}]} \tag{13}$$

Assuming $\theta_h = 1$, the probability limit of the estimated coefficient $\hat{\beta}_y^{OLS}$ is $\lambda_t \beta_y$ instead of $\beta_y$. In the case of $\lambda_t = 1$ (as implicitly assumed in the discussion of transitory fluctuations) this does no harm, but in general, the estimator will be inconsistent and the inconsistency varies as a function of age $t$ at which earnings are observed. Focusing on the impact of $\theta_h$ (setting $\lambda_t = 1$), it is not clear whether the combination of transitory fluctuations and lifecycle variation leads to an amplification bias instead of an attenuation bias. For $\lambda_h > 1$ the estimation is downward biased, but for values smaller than one and minor transitory variance the opposite is true. $\theta_h$ is a summary measure of the attenuation bias resulting from transitory fluctuations as well as lifecycle variation. Therefore the age composition of the sample matters (Jenkins 1987; Grawe 2006). In summary, neither measurement error in the offspring's status

[5] See Card (1999) for a recent survey.

[6] See Mincer (1975) and Baker (1997) among others and Vogel (2007) for an application to intergenerational mobility.

**Figure 1**
**Estimated correlation between current and permanent earnings**

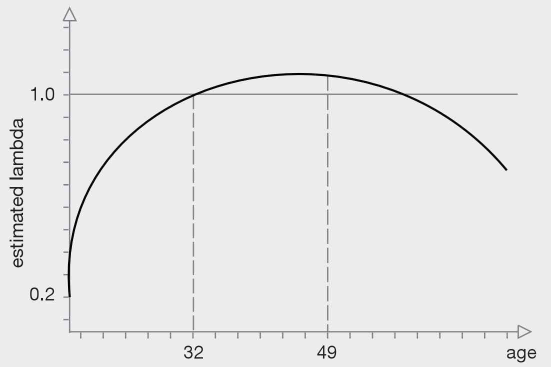
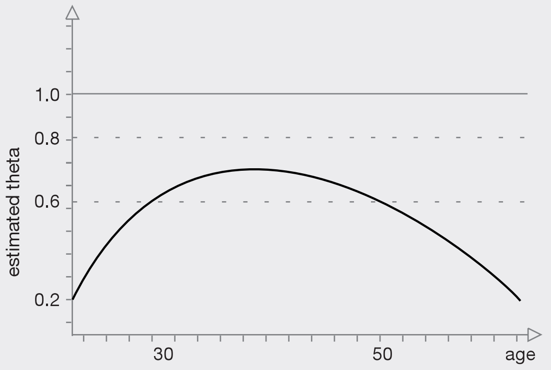



**Figure 2**
**Estimated reliability ratio**




nor measurement error in the parental long-run status is innocuous for consistency. Both induce either amplification or attenuation bias of the OLS estimation.

Using U.S. Social Security Administration earnings histories of members of the Health and Retirement Study sample, Haider & Solon (2006) assess the magnitude of measurement error in the offspring's and the parents' permanent earnings separately. Their dataset ranges from 1951 to 1991 and provides almost full-career earnings histories for a broadly representative sample of the U.S. population. This makes it possible to derive a more precise estimate of the (logarithmized) present value of lifetime earnings $\ln V^i$. Starting with the impact of measurement error in the offspring's (permanent) earnings level, the forward regression of $\ln V^i$ on $y^i_{t,h}$ leads to the estimated slope coefficient $\hat{\lambda}_{t,h}$ depicted in figure 1. Starting at a value around $\hat{\lambda}_{t,h} = 0.2$ it increases steadily. At age 32, the textbook assumption of $\lambda_{t,h} = 1$ seems reasonable. Thenceforward, $\hat{\lambda}_{t,h}$ declines some in the late forties. Turning to the case of measurement error in parental permanent earnings, the estimated reliability ratio $\hat{\theta}_h$ is depicted in figure 2. It is the result of a backward regression of $\ln V^i$ on a 5-year average of $y^i_{t,h}$. A significant increase until the age of 30 is followed by a quite robust factor between 0.6 and 0.8, but after the age of 50, $\hat{\theta}_h$ declines and the bias rises. Unfortunately, we are not aware of any comparable work for the case of Germany.

## 2.2 Sample homogeneity

In selected sub-populations with respect to location, socioeconomic status or occupation, the sample variance in long-run economic status is possibly less than in the whole population. For example, a study by Sewell & Hauser (1975) was based on a selective sample of sons from Wisconsin who graduated in 1957 and thus excluded high-school dropouts, leaving only rather successful sons in the sample. Similarly, Behrman & Taubman (1985) confined their study to parental data on white male twins born between 1927 and 1929, who both served in the army. Presumably, this father-sample is quite homogeneous. Both types of selectivity may introduce a third source of inconsistency as Solon (1989) points out. To concentrate on the effect of sample homogeneity, long-run status is assumed to be measured correctly until indicated otherwise. Formally speaking, the parent/offspring-sample is more homogeneous in long-run status, if the variance in permanent earnings $Var[y^*_{j=0,1}]$ is only a fraction $\tau$ of the population variance $Var[y_{j=0,1}]$.

$$Var[y^*_{j=0,1}] = \tau Var[y_{j=0,1}] \tag{14}$$

Under normality of parental economic status, selection on the dependent variable leads to a proportional change in the estimated intergenerational elasticity, where $R^2$ is the coefficient of determination of the population-based regression model (Goldberger 1981).

$$\text{plim } \hat{\beta}^{OLS}_{y^*} = \phi \beta_y < \beta_y \tag{15}$$

$$\phi = \frac{\tau}{1 - R^2(1-\tau)} \tag{16}$$

If $\tau < 1$ (implying $\phi < 1$) the estimated intergenerational elasticity $\hat{\beta}^{OLS}_{y^*}$ is downward inconsistent even though long-run status is measured correctly.

A sample exhibiting homogeneity in parental earnings does not affect the consistency of intergenerational elasticity estimates. This is true as long as the economic status is measured correctly. If this is not the case, the downward bias is worsened (Solon 1992; Wiegand 1997), see equation (17).

$$\left(\frac{Var[y_0^*]}{Var[y_0^*] + Var[v_{0h}]}\right)\beta_y = \text{plim } \hat{\beta}_{y^*}^{OLS} < \text{plim } \hat{\beta}_y^{OLS} = \left(\frac{Var[y_0]}{Var[y_0] + Var[v_{0h}]}\right)\beta_y \tag{17}$$

In applied empirical research, inclusion in an intergenerational dataset requires both father and son to report positive labor market earnings in the periods of interest. Presumably, in such samples $\beta_y$ is underestimated, but the use of larger representative samples eases this problem. To the best of our knowledge, however, there is no available research on the magnitude of this bias.

## 3 Econometric approach and sampling procedure

We estimate the econometric model presented in equation (18). The son's observed status $y_{1t}^i$ in year $t$ is expressed as a regression function of the father's observed status $y_{0h}^i$ in year $h$, including age controls for both (Solon 1992; Zimmerman 1992; Wiegand 1997; Vogel 2007). It is derived by incorporating age-earnings profiles into equations (2) and (3) and substituting this into the basic equation (1).

$$y_{1t}^i = \beta_0 + \beta_y y_{0h}^i + \beta_1 A_{0h}^i + \beta_2 A_{0h}^{2i} + \beta_3 A_{1t}^i + \beta_4 A_{1t}^{2i} + \omega_{1t}^i \tag{18}$$

An individual's current earnings are determined by the level of permanent earnings ($y_1^i, y_0^i$), the stage in the lifecycle [($A_t^i; A_t^{2i}$), ($A_h^i; A_h^{2i}$)], a general level of economic well-being in the corresponding generation ($\alpha_1$, $\alpha_0$), and an idiosyncratic error term ($v_{1t}^i$, $v_{0h}^i$).

$$y_{1t}^i = y_1^i + \alpha_1 + \gamma_1 A_{1t}^i + \delta_1 A_{1t}^{2i} + v_{1t}^i \tag{19}$$

$$y_{0h}^i = y_0^i + \alpha_0 + \gamma_0 A_{0h}^i + \delta_0 A_{0h}^{2i} + v_{0h}^i \tag{20}$$

The empirical part is based on samples from the German Socio-Economic Panel[7] (SOEP) from 1984 to 2006. To assure comparability of real earnings observed in different years, they are adjusted by the real GDP growth rate. Our measure of long-run economic status are real[8] monthly earnings before tax and social security deductions as reported in each cross-section of the SOEP.[9] This allows international comparison.[10] Measuring all earnings variables in their natural logarithm, we choose the intergenerational earnings elasticity as our indicator of intergenerational persistence (or mobility) and use the two terms interchangeably.

Thus, our indicator is a summary measure of personal characteristics shared by parent and offspring that are valued in the German labor market. This includes similarities in educational attainment, cognitive and noncognitive skills and personal traits (Bowles, Gintis & Osborne-Groves 2001; Bowles & Gintis 2002; Heckman 2007; Pfeiffer & Reuß 2008).

A novel feature of our study is the sampling procedure. We select pairs of fathers and sons in such a way that their earnings are observed at the closest possible stage in their lifecycles. Furthermore, the bias due to transitory fluctuations and lifecycle variation is minimized, see table 1. As a start, the self-employed, who have more volatile earnings (Baker & Solon 2003; Albarrán, Carrasco & Martínez-Granado 2007; Pfeiffer 1994), are excluded. Only the full-time employed are retained in the sample, that is individuals reporting that they worked more than 35 hours in the last week. Workers from eastern Germany are excluded as well since the possibility of mobility increased dramatically after the fall of the Berlin Wall and dynamic wage growth[11] may have changed the reliability of current earnings to reflect permanent status. To avoid sample homogeneity, only the oldest sibling is included in our baseline specification (Solon 1992). Migrants, identified by their country of origin, are dropped for our basic analysis. Migration might distort the long-run relationship between the labor market earnings of father and son due to the change of the labor market (Borjas 2006; Friedberg 2000). However, we perform a separate analysis for migrants and discuss the results. For the group of fathers, moving 5-year averages of earnings and age

[7] Consult Haisken-DeNew & Frick (2005) for further information on the dataset.

[8] Deflated by the consumer price index (base year 2000) supplied by the German Federal Statistical Office.

[9] This approach is similar to Wiegand (1997), but different from Vogel (2007), who calculates a measure of yearly earnings from monthly earnings records.

[10] See Solon (1999) for a survey on intergenerational earnings mobility. We concentrate on the persistence of labor market earnings. For research using a more inclusive measure of total economic status composed of a variety of differing types of income, earnings and monetary inheritance see Piketty (2000) and Mulligan (1997) among others. It is left for future research to construct a more inclusive measure with the SOEP, since sample size is reduced and the problem of measurement error increases.

[11] See Hunt (2002) among others.

Table 1
**Final sample overview**

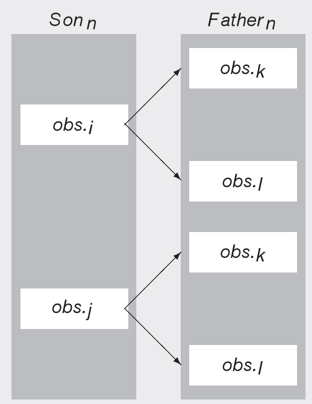
| Groups excluded from sample | Measures of economic status | Age restrictions |
|---|---|---|
| self-employed | Son | |
| part-time employed | monthly earnings (1984–2006) | |
| eastern Germans | Father | between 30–50 |
| migrants | monthly earnings (1984–2006) | |
| younger brothers | years of education | |

are calculated to reduce the attenuation bias. Thus, if for a given observation earnings are not observable in each of the four following years, it is dropped.

Furthermore, the following age restrictions are imposed (and relaxed again for further discussion) to take into account the time-varying association between short- and long-run economic status. This procedure takes into account the pattern of the variance of the transitory component over the lifecycle (which flattens out at mid age in the U.S. (Baker & Solon 2003)). Since the association between monthly and lifetime earnings is still low for workers below the age of 30, we select workers above that age. For younger workers job mobility is high and earnings are more volatile, partly because of lower tenure (Haider & Solon 2006; Björklund 1993). Workers aged over 50 are excluded as well. Labor market status and hours worked may become more volatile again, which might depress the estimated level of persistence (Grawe 2006). However, this line of reasoning may differ between countries, for instance as a result of different industrial structures or different degrees of employment protection (Blau & Kahn 1996; OECD 1999; Pries & Rogerson 2005).

Finally, father ($obs_k$ and $obs_l$) and son ($obs_j$ and $obs_i$) observations (of family $n$) satisfying the sampling rule are matched in all possible combinations, see figure 3. This procedure leads to numerous matched observations for each father-son pair. To identify a unique pair, intended to lead to the most reliable estimate of the intergenerational elasticity, a decision rule is implemented. For each observation we select the one with the smallest absolute age difference between father and son. This is to ensure that father and son are observed at the most similar stages in their lifecycle possible. If more than one observation for a particular father-son pair still fulfills the requirement, the one associated with the lowest father age is used. For comparison and discussion, other samples with less restrictive selection rules are utilized in the next section.

Figure 3
**Sampling strategy**




The sample contains 180 father-son pairs compared to Wiegand's (1997) 130 and Vogel's (2007) 300. Table 2 depicts the basic statistics. The age difference between father and son amounts to 8.68 years. Sons in the sample report lower earnings than their matched fathers which is mainly explained by the

Table 2
**Final sample statistics**

| Statistic | Fathers | Sons |
|---|---|---|
| Gross earnings in Euros[1] | 2,307.03 | 1,936.87 |
| Sd. of gross earnings | 716.82 | 640.40 |
| Year of observation | 1,987 | 2,004 |
| Age in years | 44.40 | 35.73 |
| Age difference in years | 8.68 | |
| Number of observations | 180 | |

[1] *reported 5/1 – year average of adjusted real gross monthly earnings*

early stage in their lifecycle. While most information on the father's economic status is obtained within the early SOEP waves, the collection of information regarding the offspring is not confined to the most recent wave. The age composition of our sample differs substantially from previous studies. Sons are 35 years old, which is an increase of 4 years compared to Wiegand (1997) and 13 years compared to Couch & Dunn (1997). Solon (1992) reports an average age of 29 for sons, while Björklund & Jäntti (1997) rely on sons at the age of 34 on average. An average age of 44 for fathers is slightly lower than that reported by Wiegand (1997) with 46 years, while Couch & Dunn's (1997) fathers are 51 years old. Solon's (1992) fathers are reported to be 42 years of age on average, nearly identical to an average father in Björklund & Jäntti's (1997) sample (43 years).

Selection could rise from the blind eye on individuals not meeting the selection rules, table 2. The final sample is compared to all workers meeting the sample requirements except for the need to report positive earnings 5 years in a row and being matched with their offspring. The father-sample is contrasted in 1984, while the son-sample is compared in 2004. Earnings in the father-sample are almost identical to those reported by all workers in 1984 (€ 2,331.01). However, the standard deviation is higher in the comparison group (€ 782.43) in 1984. Using 5-year averages of earnings in the father-sample, therefore, as intended, reduces transitory fluctuations. Comparing the son-sample, earnings are higher (€ 1,917.13 in the comparison group) and show a higher standard deviation (€ 574.15). In our son-sample the average age is lower, which induces greater wage dispersion.

## 4 Econometric findings

### 4.1 Basic results

The OLS estimate based on a 5-year average of earnings $\hat{\beta}_y^{OLS} = 0.282$ is higher than that obtained by Wiegand (1997), whereas the one-year snapshot is about the same, table 3. Compared to Vogel (2007), the result based on the 5-year average of earnings is similar. We use years of education[12] as

Table 3
**Basic results**

| | 5-year average earnings | Single-year earnings |
|---|---|---|
| **OLS estimate** | | |
| Intergenerational elasticity | 0.282 | 0.205 |
| 95 % confidence interval | (0.09–0.44) | (0.08–0.32) |
| Standard error | 0.087 | 0.061 |
| Observations | 180 | 249 |
| **IV estimate**[1] | | |
| Intergenerational elasticity | 0.374 | |
| 95 % confidence interval | (0.09–0.65) | |
| Standard error | 0.144 | |
| Observations | 180 | |

[1] *using years of education*

an instrument to bracket the intergenerational elasticity. According to the IV estimate the intergenerational elasticity is higher, $\hat{\beta}_y^{IV} = 0.374$. Following Solon's (1992) approach, the intergenerational elasticity of German workers should lie between the two estimates and we suggest a reasonable value of 1/3.

$$\boxed{\hat{\beta}_y^{OLS} = 0.282 < \beta_y < 0.374 = \hat{\beta}_y^{IV}}$$

The 95 % confidence interval of the IV estimate $[0.09 \leq \hat{\beta}_y^{IV} \leq 0.66]$ includes the OLS estimate. Although the two point estimates contain some useful information, the degree of precision seems to be rather low. We come back to this issue in the conclusion.

### 4.2 Investigating the bias from transitory fluctuations

Tables 4 and 5 report the general pattern that $\hat{\beta}_y^{OLS}$ increases with the number of years averaged as the attenuation bias declines. This is in line with equation (6). For inclusion in the balanced panel, fathers' earnings need to be observed for 5 years in a row even though only lower averages are used for the supplementary estimations. The changing estimate is due to the reduced number of years averaged and not to a change in the sample composition. For this

[12] This variable includes both school and occupational education. The German school system introduces differentiated educational tracks after four grades of primary education. The lower secondary school (Hauptschule) graduates individuals after five years of secondary education and is traditionally a preparation for blue-collar occupations. The intermediate secondary school (Realschule) lasts six years and prepares pupils for white-collar employment. The highest track (Gymnasium) offers nine years of schooling and a qualification (Abitur) which provides access to academic studies. Completion of an apprenticeship adds another 1.5 years, a technical college 3 years, and graduation from university increases years of education by 5 years.

Table 4

**Summary results[1]: balanced panel**

| Father measure[2] | 5–year | 4–year | 3–year | 2–year | 1–year |
|---|---|---|---|---|---|
| Intergenerational elasticity | 0.2822*** | 0.2841*** | 0.2751*** | 0.2441*** | 0.1984** |
| 95 % confidence interval | (0.09–0.44) | (0.11–0.45) | (0.11–0.44) | (0.08–0.41) | (0.04–0.35) |
| Standard error | 0.0870 | 0.0866 | 0.0854 | 0.0841 | 0.0798 |
| Observations | 180 | 180 | 180 | 180 | 180 |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
[1] *see Table 10 in the Appendix for the detailed results*
[2] *average of fathers' logarithmized adjusted real gross monthly earnings*

Table 5

**Summary results[1]: unbalanced panel**

| Father measure[2] | 5–year | 4–year | 3–year | 2–year | 1–year |
|---|---|---|---|---|---|
| Intergenerational elasticity | 0.2822*** | 0.2867*** | 0.2596*** | 0.2076*** | 0.2045*** |
| 95 % confidence interval | (0.11–0.45) | (0.13–0.45) | (0.10–0.41) | (0.07–0.34) | (0.08–0.32) |
| Standard error | 0.0870 | 0.0815 | 0.0790 | 0.0695 | 0.0614 |
| Observations | 180 | 190 | 217 | 227 | 249 |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
[1] *see Table 11 in the Appendix for the detailed results*
[2] *average of fathers' logarithmized adjusted real gross monthly earnings*

reason, the number of observations remains constant. The unbalanced panel, however, includes all pairs with the necessary number of successive earnings observations for the father that is needed for the respective estimation. A comparison of the OLS results in the balanced and unbalanced panels reveals that the difference between a 5-year and a 4-year average of fathers' earnings is negligible. However, it makes a difference in our sample whether the estimate is based on a 1/2-year average or a 4/5-year average. Averaging only a small number of years amplifies the attenuation bias due to a high volatility of the earnings measure utilized. This result is in line with the literature as reported in section 2.

The rather early decrease of the estimated coefficient in the unbalanced panel might be attributable to the construction of the panel. When lowering the number of years averaged, the added individuals do not report earnings in the following year probably due to unemployment or part-time employment. This implies that father-son pairs with larger transitory fluctuations are added to the panel consecutively.

## 4.3 Investigating the bias from lifecycle variation

Raising the upper age limit from 50 to 55 results in a rather sharp increase in sample size and a slight decrease in estimated intergenerational persistence. However, table 6 reveals an increase in the estimate when continuing to relax the age restriction. This seems to be in line with Vogel (2007), whose estimate of intergenerational persistence in Germany included individuals aged over 50 and is slightly higher. We offer two explanations. First, the increase could point to sample selection with only pairs added that exhibit a particularly strong persistence of earnings. However, a comparison of the descriptive statistics (years of education, monthly earnings) did not provide any evidence on the type of selection. Second, the increase in the estimated level

Table 6

**Summary results[1]: relaxing age restrictions of fathers**

| Fathers' maximum age | 50 | 55 | 60 | 65 |
|---|---|---|---|---|
| Intergenerational elasticity[2] | 0.2822*** | 0.2509*** | 0.3538*** | 0.3584*** |
| 95 % confidence interval | (0.11–0.45) | (0.10–0.41) | (0.22–0.49) | (0.22–0.49) |
| Standard error | 0.0870 | 0.0794 | 0.0696 | 0.0686 |
| observations | 180 | 240 | 281 | 285 |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] see Table 13 in the Appendix for the detailed results*
*[2] 5-year average of fathers' logarithmized adjusted real gross monthly earnings; son at least 30 years of age*

Table 7

**Summary results[1]: relaxing age restrictions for sons**

| Sons' minimum age | 30 | 25 | 20 |
|---|---|---|---|
| Intergenerational elasticity[2] | 0.2822*** | 0.2553*** | 0.2402*** |
| 95 % confidence interval | (0.11–0.45) | (0.12–0.39) | (0.13–0.30) |
| Standard error | 0.0870 | 0.0666 | 0.0558 |
| Observations | 180 | 282 | 385 |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] see Table 14 in the Appendix for the detailed results*
*[2] 5-year average of fathers' logarithmized adjusted real gross monthly earnings; father at most 50 years of age*

of mobility could be explained by an increase in the reliability ratio $\theta_h$ in our sample rather than a decrease as documented for the United States, figure 2. This is presumably the result of the comparatively high degree of centralization governing wage determination in Germany and employment protection laws which, together with the accumulation of specific human capital, favor incumbent workers.[13] This could reduce the transitory fluctuations among older German workers.

Table 7 documents a significant rise in the number of observations and a sharp decline in the estimated intergenerational elasticity when the age requirement for sons is consecutively lowered to 20 years. This seems to be in line with Haider & Solon (2006). The parameter $\lambda_t$ (see equation (12) in section 2) is lowered as younger and younger workers are added to the sample and the lifecycle bias rises.

The analysis above gave the impression that the age composition of either sample is changed without affecting the other. Obviously, this is not true since father-son *pairs* are added. However, negligible changes in the age composition of the unchanged sample (with respect to the age restrictions imposed) support this approach.

## 4.4 An analysis of migrants and further sensitivity checks

### *Analysis of migrants*

The analysis of the migrant population[14] is based on 93 father-son pairs when relying on a 5-year average of earnings, see table 15 in the Appendix for the detailed results. It emerges that the point estimate

[13] See Botero, Djankov, Porta & Lopez-De-Silanes (2004) and Franz & Pfeiffer (2006) among others.

[14] Identified by the fact that at least the father was not born in Germany.

of the intergenerational earnings elasticity ($\hat{\beta}_y^{OLS}$ = 0.41) is higher compared to our sample of German workers. While the age structure of the samples are fairly identical, there are substantial differences in average earnings. For migrants, sons' earnings are lower (€ 1,617 compared to € 1,936). The same is true for migrant fathers, who earn € 1,789 on average compared to € 2,307 for a German father. This suggests possible non-linearities in the intergenerational link. For example Corak & Heisz (1999) and Hertz (2005) present evidence of stronger persistence among low-income families.

### Including younger siblings

The inclusion of younger siblings raises the sample size from 180 to 224 when relying on a 5-year average of fathers' earnings. The point estimate is slightly reduced to $\hat{\beta}_y^{OLS} = 0.276$. Siblings share the same family and community background, which makes similar long-run economic status more likely and increases homogeneity within the sample. This depresses the estimated coefficient slightly, see table 16 in the Appendix for the detailed results.

### Adjustment of monthly earnings

To ensure robustness with respect to the measure of comparability (GDP growth in the baseline estimation), earnings are deflated by the growth rate of average real gross monthly earnings in Germany's manufacturing sector (as reported by the German Federal Statistical Office). The estimated intergenerational elasticity is not affected, see table 17 in the Appendix for the detailed results.

### Instrumenting parental status

To compare our findings with Wiegand (1997), the IV estimation is repeated instrumenting parental status using the Wegener index, a standard index for occupational prestige. The baseline estimate ($\hat{\beta}_y^{IV}$ = 0.372) remains unchanged. The finding that both instruments lead to fairly identical results is robust to changes in the sampling rule, see table 18 in the Appendix for the detailed results.

## 5 Discussion and concluding remarks

Table 8 compares our result to the international evidence. Although the studies differ with respect to data and methods, the comparison suggests higher mobility (that is less persistence) in Germany compared to the United States and the United Kingdom, but lower mobility than Sweden.

Table 8
**International perspective**

| Country | OLS result | IV result |
|---|---|---|
| United States | | |
| Mazumder (2005) | 0.613 (0.09)[a] | – |
| Solon (1992) | 0.413 (0.09) | 0.526 (0.14) |
| Zimmerman (1992) | 0.400 (0.06) | 0.330 (0.27) |
| United Kingdom | | |
| Dearden, Machin & Reed (1997) | 0.240 (0.03) | 0.443 (0.03) |
| Sweden | | |
| Björklund & Jäntti (1997) | 0.216 (0.04) | – |
| Germany | | |
| Couch & Dunn (1997) | 0.124 (0.07) | – |
| Wiegand (1997) | 0.238 (0.06) | 0.402 (0.13) |
| Vogel (2007) | 0.266 (0.06) | – |
| this paper | 0.282 (0.09) | 0.374 (0.14) |

[a] *standard errors in parentheses*

Our preferred point estimate of the elasticity in Germany is $\beta_y^{GER} = 1/3$, compared with $\beta_y^{US} = 0.4$ for the United States and $\beta_y^{S} = 0.2$ for Sweden. In comparison to former studies by Couch & Dunn (1997) and Wiegand (1997) on intergenerational persistence in Germany, Vogels' (2007) and ours suggest higher persistence. This is the result of our special attention to the sources of potential lifecycle bias. However, what is common to all studies presented in table 8 are the considerable confidence intervals, which currently prevent any strong comparative statements on the level of intergenerational persistence.

For illustrative purposes, we conclude by working out some consequences of the value of $\beta_y^{GER} = 1/3$ for Germany. The intergenerational elasticity $\beta_y$ translates intragenerational inequality in parental long-run labor market status into the economic advantage which a child of parents with a higher economic status can hope for compared to a child of lower-status parents. Table 9 depicts the advantage of a child with parents in the top permanent earnings decile compared to offspring born to parents in the bottom decile as determined by equation (21) (Corak 2004).

$$\frac{y_1^{90th}}{y_1^{10th}} = \left(\frac{y_0^{90th}}{y_0^{10th}}\right)^{\beta_y} \tag{21}$$

For Germany, Gernandt & Pfeiffer (2007) calculate a 90/10-percentile earnings ratio of 2.5 for a cross-

Table 9
**Inequality and the expected permanent earnings advantage**

| | Intergenerational elasticity | | | |
|---|---|---|---|---|
| **90/10 – ratio** | **0.2** | **1/3** | **0.4** | **0.5** |
| 2.0 | 15 % | 25 % | 32 % | 41 % |
| 2.5 | 20 % | 35 % | 44 % | 58 % |
| 3.0 | 25 % | 44 % | 55 % | 73 % |
| 3.5 | 28 % | 51 % | 65 % | 87 % |
| 4.0 | 32 % | 59 % | 74 % | 100 % |

section sample of prime-age dependent male workers in 2005, which is quite similar to our sample. Then, taking our advocated value for an intergenerational elasticity in Germany of $\beta_y^{GER} = 1/3$, the expected earnings advantage amounts to 35 %. If $\beta_y^{GER} = 1/3$ were 0.5, the advantage would increase to 59 %. We would like to add that a value of $0 < \beta_y < 1$ does not necessarily imply a compression of the earnings distribution, because the variance of the error term $\varepsilon_1^i$ in equation (1) matters, too.

Summarizing our findings, we find that intergenerational earnings persistence among western German workers is higher than previously suggested. A value of $\beta_y^{GER} = 1/3$ still indicates that there is substantial intergenerational mobility, which is presumably a result of the massive expansion of publicly funded education in Germany from the seventies onwards and the openness of the German labor market.

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## Estimation appendix

Table 10
### Detailed results: balanced panel

| | 5-year average | | 4-year average | | 3-year average | | 2-year average | | 1-year average | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | | | |
| Measure[2] | 0.2822*** | 0.0870 | 0.2841*** | 0.0866 | 0.2751*** | 0.0854 | 0.2441*** | 0.0841 | 0.1984** | 0.0798 |
| Age | –0.1100 | 0.0735 | –0.1106 | 0.0734 | –0.1072 | 0.0733 | –0.1040 | 0.0738 | –0.0968 | 0.0741 |
| Age (squared) | 0.0015* | 0.0009 | 0.0016* | 0.0009 | 0.0015* | 0.0009 | 0.0015 | 0.0009 | 0.0014 | 0.0009 |
| Son | | | | | | | | | | |
| Age | 0.2245** | 0.1050 | 0.2249** | 0.1049 | 0.2249** | 0.1050 | 0.2160** | 0.1054 | 0.2150** | 0.1061 |
| Age (squared) | –0.0029** | 0.0014 | –0.0029** | 0.0014 | –0.0029** | 0.0014 | –0.0028* | 0.0014 | –0.0028* | 0.0014 |
| F - Test | 5.6816*** | | 5.7364*** | | 5.6521*** | | 5.2225*** | | 4.7309*** | |
| Adjusted $R^2$ | 0.1156 | | 0.1168 | | 0.1150 | | 0.1055 | | 0.0944 | |
| Observations | 180 | | 180 | | 180 | | 180 | | 180 | |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
[1] *Coeff. = coefficient; SE. = standard error*
[2] *average of fathers' logarithmized adjusted real gross monthly earnings*

Table 11
### Detailed results: unbalanced panel

| | 5-year average | | 4-year average | | 3-year average | | 2-year average | | 1-year average | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | | | |
| Measure[2] | 0.2822*** | 0.0870 | 0.2867*** | 0.0815 | 0.2596*** | 0.0790 | 0.2076*** | 0.0695 | 0.2045*** | 0.0614 |
| Age | –0.1100 | 0.0735 | –0.0983 | 0.0754 | –0.0725 | 0.0730 | –0.0553 | 0.0744 | –0.1027 | 0.0687 |
| Age (squared) | 0.0015* | 0.0009 | 0.0014 | 0.0009 | 0.0010 | 0.0009 | 0.0008 | 0.0009 | 0.0014* | 0.0008 |
| Son | | | | | | | | | | |
| Age | 0.2245** | 0.1050 | 0.2333** | 0.1033 | 0.2433** | 0.1011 | 0.2384** | 0.0990 | 0.2647*** | 0.0904 |
| Age (squared) | –0.0029** | 0.0014 | –0.0030** | 0.0014 | –0.0031** | 0.0014 | –0.0031** | 0.0013 | –0.0034*** | 0.0012 |
| F - Test | 5.6816*** | | 6.5795*** | | 6.1001*** | | 5.9149*** | | 7.1282*** | |
| Adjusted $R^2$ | 0.1156 | | 0.1286 | | 0.1056 | | 0.0982 | | 0.1100 | |
| Observations | 180 | | 190 | | 217 | | 227 | | 249 | |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
[1] *Coeff. = coefficient; SE. = standard error*
[2] *average of fathers' logarithmized adjusted real gross monthly earnings*

Table 12

## Detailed results: IV – estimation using years of education

| | Lower age limit –5 | | Basic specification | | Upper age limit +5 | |
|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | |
| Measure[2] | 0.3241*** | 0.1139 | 0.3735** | 0.1435 | 0.3365*** | 0.1269 |
| Son | | | | | | |
| Age | 0.2149*** | 0.0478 | 0.1966* | 0.1049 | 0.2290*** | 0.0828 |
| Age (squared) | –0.0027*** | 0.0007 | –0.0025* | 0.0014 | –0.0029*** | 0.0011 |
| F - Test | 34.6628*** | | 4.6729*** | | 5.7988*** | |
| Adjusted $R^2$ | 0.2693 | | 0.0717 | | 0.0752 | |
| Observations | 282 | | 180 | | 240 | |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] years of education*

Table 13

## Detailed results: relaxing age restrictions for fathers

| | Max. age 50 | | Max. age 55 | | Max. age 60 | | Max. age 65 | |
|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | |
| Measure[2] | 0.2822*** | 0.0870 | 0.2509*** | 0.0794 | 0.3538*** | 0.0696 | 0.3584*** | 0.0686 |
| Age | –0.1100 | 0.0735 | 0.0909* | 0.0524 | 0.0488 | 0.0378 | 0.0536 | 0.0341 |
| Age (squared) | 0.0015* | 0.0009 | –0.0010* | 0.0006 | –0.0005 | 0.0004 | –0.0006 | 0.0004 |
| Son[3] | | | | | | | | |
| Age | 0.2245** | 0.1050 | 0.2008** | 0.0850 | 0.1377** | 0.0635 | 0.1258** | 0.0618 |
| Age (squared) | –0.0029** | 0.0014 | –0.0026** | 0.0011 | –0.0017** | 0.0008 | –0.0016* | 0.0008 |
| F - Test | 5.6816*** | | 5.3447*** | | 7.5867*** | | 7.8214*** | |
| Adjusted $R^2$ | 0.1156 | | 0.0833 | | 0.1052 | | 0.1072 | |
| Observations | 180 | | 240 | | 281 | | 285 | |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] 5-year average of fathers' logarithmized adjusted real gross monthly earnings*
*[3] son at least 30 years of age*

Table 14

## Detailed results: relaxing age restrictions for sons

| | Min. age 30 | | Min. age 25 | | Min. age 20 | |
|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. |
| Father[3] | | | | | | |
| Measure[2] | 0.2822*** | 0.0870 | 0.2553*** | 0.0666 | 0.2402*** | 0.0558 |
| Age | –0.1100 | 0.0735 | –0.1139** | 0.0477 | –0.0806** | 0.0373 |
| Age (squared) | 0.0015* | 0.0009 | 0.0015** | 0.0006 | 0.0011** | 0.0005 |
| Son | | | | | | |
| Age | 0.2245** | 0.1050 | 0.2275*** | 0.0465 | 0.1227*** | 0.0248 |
| Age (squared) | –0.0029** | 0.0014 | –0.0029*** | 0.0007 | –0.0015*** | 0.0004 |
| F - Test | 5.6816*** | | 24.2022*** | | 48.0827*** | |
| Adjusted R² | 0.1156 | | 0.2922 | | 0.3801 | |
| Observations | 180 | | 282 | | 385 | |

*Basic specification*
*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] 5-year average of fathers' logarithmized adjusted real gross monthly earnings*
*[3] father at most 50 years of age*

Table 15

## Robustness check I: analysis of migrant sample

| | 5-year average | | 4-year average | | 3-year average | | 2-year average | | 1-year average | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | | | |
| Measure[2] | 0.4056*** | 0.1362 | 0.3598*** | 0.1326 | 0.3183** | 0.1334 | 0.2808** | 0.1241 | 0.2732** | 0.1073 |
| Age | 0.2708 | 0.1518 | 0.2134 | 0.1466 | 0.1507 | 0.1233 | 0.1300 | 0.1192 | 0.0026* | 0.0953 |
| Age (squared) | –0.0033 | 0.0019 | –0.0026 | 0.0018 | –0.0018 | 0.0015 | –0.0030 | 0.0014 | –0.0033* | 0.0011 |
| Son | | | | | | | | | | |
| Age | 0.2556* | 0.1411 | 0.2437* | 0.1377 | 0.1949* | 0.1128 | 0.2267** | 0.1100 | 0.2442** | 0.1076 |
| Age (squared) | –0.0032* | 0.0019 | –0.0030 | 0.0019 | –0.0024 | 0.0015 | –0.0028* | 0.0015 | –0.0031** | 0.0015 |
| F - Test | 4.26*** | | 4.08*** | | 3.27*** | | 4.00*** | | 4.20*** | |
| Adjusted R² | 0.1504 | | 0.1372 | | 0.1275 | | 0.1042 | | 0.0981 | |
| Observations | 93 | | 98 | | 118 | | 130 | | 148 | |

*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] average of fathers' logarithmized adjusted real gross monthly earnings*

Table 16

## Robustness check II: including younger siblings

| | 5-year average | | 4-year average | | 3-year average | | 2-year average | | 1-year average | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | | | |
| Measure[2] | 0.2767*** | 0.0833 | 0.2726*** | 0.0831 | 0.2639*** | 0.0827 | 0.2300*** | 0.0820 | 0.1780** | 0.0784 |
| Age | -0.1315* | 0.0722 | -0.1308* | 0.0722 | -0.1289* | 0.0722 | -0.1255* | 0.0726 | -0.1185 | 0.0730 |
| Age (squared) | 0.0018** | 0.0009 | 0.0018* | 0.0009 | 0.0018* | 0.0009 | 0.0017* | 0.0009 | 0.0016* | 0.0009 |
| Son | | | | | | | | | | |
| Age | 0.2518** | 0.0992 | 0.2508** | 0.0993 | 0.2509** | 0.0994 | 0.2452** | 0.0999 | 0.2440** | 0.1005 |
| Age (squared) | -0.0032** | 0.0014 | -0.0032** | 0.0014 | -0.0032** | 0.0014 | -0.0032** | 0.0014 | -0.0031** | 0.0014 |
| F - Test | 6.7625*** | | 6.7032*** | | 6.5778*** | | 6.0692*** | | 5.4723*** | |
| Adjusted R² | 0.1144 | | 0.1134 | | 0.1112 | | 0.1021 | | 0.0911 | |
| Observations | 224 | | 224 | | 224 | | 224 | | 224 | |

*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] average of fathers' logarithmized adjusted real gross monthly earnings*

Table 17

## Robustness check III: earnings adjusted using growth rate of wages in manufacturing

| | 5-year average | | 4-year average | | 3-year average | | 2-year average | | 1-year average | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | | | | | |
| Measure[2] | 0.2806*** | 0.0873 | 0.2816*** | 0.0869 | 0.2720*** | 0.0857 | 0.2414*** | 0.0845 | 0.1976** | 0.0803 |
| Age | –0.1072 | 0.0734 | –0.1078 | 0.0733 | –0.1048 | 0.0733 | –0.1021 | 0.0737 | –0.0955 | 0.0741 |
| Age (squared) | 0.0015 | 0.0009 | 0.0015 | 0.0009 | 0.0015 | 0.0009 | 0.0014 | 0.0009 | 0.0014 | 0.0009 |
| Son | | | | | | | | | | |
| Age | 0.2396** | 0.1050 | 0.2399** | 0.1050 | 0.2395** | 0.1051 | 0.2308** | 0.1055 | 0.2298** | 0.1062 |
| Age (squared) | –0.0031** | 0.0014 | –0.0031** | 0.0014 | –0.0031** | 0.0014 | –0.0030** | 0.0014 | –0.0030** | 0.0014 |
| F - Test | 5.7383*** | | 5.7774*** | | 5.6818*** | | 5.2625*** | | 4.7995*** | |
| Adjusted R² | 0.1169 | | 0.1177 | | 0.1157 | | 0.1064 | | 0.0959 | |
| Observations | 180 | | 180 | | 180 | | 180 | | 180 | |

*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] average of fathers' logarithmized adjusted real gross monthly earnings*

Table 18

**Robustness check IV: IV – estimation using Wegener index**

| | Lower age limit –5 | | 30–50 age limit | | Upper age limit +5 | |
|---|---|---|---|---|---|---|
| | Coeff.[1] | SE.[1] | Coeff. | SE. | Coeff. | SE. |
| Father | | | | | | |
| Measure[2] | 0.3809*** | 0.1116 | 0.3725*** | 0.1421 | 0.3223*** | 0.1219 |
| Son | | | | | | |
| Age | 0.2029*** | 0.0480 | 0.2121* | 0.1077 | 0.2421*** | 0.0851 |
| Age (squared) | -0.0026*** | 0.0007 | -0.0027* | 0.0015 | -0.0031*** | 0.0011 |
| F - Test | 34.3379*** | | 4.7035*** | | 5.9378*** | |
| Adjusted $R^2$ | 0.2607 | | 0.0717 | | 0.0752 | |
| Observations | 273 | | 174 | | 231 | |

*Source: own calculations*
*Level of significance: *** 1 % ** 5 % * 10 %*
*[1] Coeff. = coefficient; SE. = standard error*
*[2] Wegener index of occupational prestige*